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LABOR MARKET CHARACTERISTICS AND THE LABOR FORCE
PARTICIPATION OF INDIVIDUALS

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LABOR MARKET CHARACTERISTICS AND THE LABOR FORCE PARTICIPATION OF INDIVIDUALS

If there is a distinctively sociological perspective on human behavior, it is one that is highly sensitive to the way that seemingly individual-level processes are shaped and conditioned by the cultural (e.g., Durkheim, 1895) or material (e.g., Marx and Engels, 1922) conditions of the collectivities in which people live. In its most distilled form, this perspective focuses neither on the properties or collectivities nor on the characteristics of individuals, but on the interaction between individual and collective phenomena in shaping the actions of individuals. In this paper, we apply this perspective to the analysis of married women's labor force participation. At the theoretical level, we formulate a set of hypotheses about the ways in which demographic and economic characteristics of labor markets intensify or weaken the impact of married women's individual characteristics on their individual probabilities of labor force participation. At the empirical level, we test some of those hypotheses with a combination of individual and aggregate level probit and regression analyses of data from 409 distinct geographic areas. These analyses allow us to measure interactions among individual [illegible] aggregate level determinants of these interactions. We begin by formu[illegible]ne problem which we investigate here, and then move on to specific hypotheses, methodological considerations and data analyses. We conclude with a discussion of theoretical and applied consequences of our findings.

Labor Force Participation and the Constraints of Motherhood

It is a commonplace of sociology, demography, economics and social policy that in industrial societies the social organization of families, and of women's lives in general, is deeply affected by the constraints which childbearing and childrearing place on women's labor force participation. In empirical studies, the presence of children, their numbers and their ages are standard, and usually dominant, determinants of whether or not a wife works for pay or seeks to do so (Sweet, 1973; Cain, 1966; Heckman and Willis, 1977). Although there is evidence that long-run decisionmaking about childbearing and long-range planning for paid employment involves reciprocal causation (see, e.g., Waite and Stolzenberg, 1976), there is ample reason to believe that in the short run, the constraints of actual childbearing and childrearing on actual labor force participation are direct and strong. In particular, the irreversibility of a birth, and the immediacy of childcare demands would seem to produce a powerful effect of childbearing and childrearing on whether or not a mother works for pay during the period following childbirth.[1] Consistent with this reasoning, there is evidence that virtually all types of wives, even those who plan long-term employment, tend to reduce their labor force participation shortly before and after a birth (Hout, 1978; Cramer, 1980; Shapiro and Mott, 1979). Further, there is evidence that mothers' efforts to enter or re-enter the labor force are, cn average, sharply constrained by their ability to find

[1]Obviously, current labor force participation by a woman cannot alter the number or ages of children that she has borne already, and so the effect of current labor force participation on a woman's present family size is nil.

acceptable substitutes for their own childcare activities (Shortlidge, 1975; Presser and Baldwin, 1980).[2] This evidence would seem consistent with the inference that the constraint of children is more severe in places where childcare is more expensive or difficult to arrange than in places where childcare is less expensive or difficult to arrange. However, this inference has not been tested. Indeed, there has been no systematic effort to measure or model the way that *aggregate properties* of communities or labor markets affect the impact of women's individual-level characteristics on their labor force participation.

Perhaps the best summary of knowledge about the relationship between individual and aggregate-level determinants of labor force participation is given in Sweet's (1973:23) well-known work on female labor force activity.

> We will not consider any...labor market characteristics in this study. Their interrelation with the individual characteristics is undoubtedly very complex. We will assume that they do not have effects that would modify the effects of the variables we will consider... This is an important area of study which has not yet been dealt with at all.

Apparently with reasoning similar to Sweet's, Cohen et al. (1970) include both individual and aggregate-level variables in their models of women's labor force participation, but do not allow for interactions between labor market variables and individual variables. But Cohen et al. do test for additive effects of several area characteristics, including per capita personal income in the woman's city (SMSA) of

[2]This number differs substantially by race; Shortlidge (1975) reports that nearly 50 percent of black mothers and 10 percent of white mothers surveyed responded that they are kept out of the job market because of lack of acceptable child care.

residence, the unemployment rate in that area, recent percentage changes in employment levels and the gender composition of the city's labor force. Of these, only per capita personal income fails to affect the probability that individual wives participate in the labor force. Similarly, in a study of Canadian data, Nakamura et al. (1979) conclude that individual and family decisionmaking and aggregate-level labor market conditions both play important roles in wives' labor supply decisions. Morgan et al. (1966) find that the size of the town in which individual women live affects their probability of labor force participation. And well-known aggregate-level studies by Bowen and Finegan (1969) find that plentiful job opportunities, relatively high wages, low labor market competition, and cheap household help raise the labor force participation rate of wives. But none of these studies allow for the possibility that aggregate conditions affect the linkage between wives' individual characteristics (such as the number and ages of their children) and their probability of labor force participation.[3]

We now suggest that there are straightforward reasons to hypothesize that aggregate-level characteristics of the places (labor markets) in which women live do significantly affect the impact of the number and ages of a wife's children on her probability of labor force participation. That is, we argue that there are powerful interactions between aggregate characteristics of places where women live and some of

[3]With the exception of Nakamura (1979) all of these studies were done in the 1960s or used data from the 1960s, a period when few mothers of young children worked for pay. Subsequently there have been large increases in the mean and variance of the probability that mothers of young children participate in the labor force.

the most important individual-level determinants of their probability of labor force participation. In making this argument, our reasoning is based on the observation (illustrated below) that the mechanisms which link a woman's childrearing responsibilities to her labor force participation generally involve phenomena which are characteristics of the city or community in which the woman lives. Where these city or community characteristics vary, we argue, the mechanisms which are based on them will also tend to vary. The hypotheses presented below will illustrate this point.

Hypotheses

We hypothesize that the greater the availability of child care in a local area, the smaller the constraint of children on the labor market activity of wives. In addition, we expect that this effect of childcare availability decreases with the increasing age of the woman's child(ren). We expect these relationships because time spent in caring for children constitutes the most important, time-intensive and employment-inhibiting household responsibility of wives with young children. Someone must care for a preschool-aged child for the entire day; if a family cannot find another person to perform this task then one of the parents must do it and that parent is usually the mother. Of course, not every family with a working wife pays someone else to care for its children. In fact, in most cases family members provide some daytime care of the young children of an employed mother. Fathers and older siblings watch younger children. And some employed mothers supervise their children while they work (U.S. Bureau of the Census,

1976). But a very substantial minority of all young children with employed mothers receive care in day care centers, in their own home from a nonrelative or from a nonrelative in that person's home. More precisely, in 1977, 40 percent of all employed mothers with children under five years old used one of these types of non-familial childcare while they worked (Presser and Baldwin, 1980:1209). And even children who do not receive daily care from someone outside their immediate family probably receive such care occasionally. So, the availability of child care would seem to be a major factor in releasing mothers from the child supervision responsibilities which would keep them home from the labor market. Thus, we expect that the greater the ease with which families make childcare arrangements--regularly or occasionally--the greater the ease with which the mother of young children can work and the smaller the constraint of children on the mother's probability of labor force participation. Since the availability of child care would seem to be a major determinant of the ease of arranging child care, we hypothesize that _the greater the availability of child care in a local area, the smaller the constraints of children on the market activity of wives who live there_.

Obviously, the need for child care varies with the age of the child. For infants, care is continuous and the constraints of its demands are virtually total. For slightly older children, school provides needed supervision for much of the day. And older children are able to fend for themselves in an employed mother's absence. Accordingly, we expect (as have others, e.g., Sweet, 1973) that the constraint of children on their mother's labor force participation

decreases as the children grow older. Since childcare duties become less rigid as children grow older, we also expect (as others have not expected) that the effect of childcare availability on the constraint of children is weaker for older children than for younger ones.

In addition, we suspect that the labor force participation of married mothers depends on the cost of child care as well as its availability. Other things equal, the higher the cost of child care in an area, the fewer the wives for whom employment provides a net financial gain over full-time housewifery. Consistent with this reasoning, Hofferth (1979:652) points out that the cost of care is one of the most important determinants of the type of childcare arrangements that parents make. In aggregate-level analyses, Bowen and Finegan (1969) find that median wages of domestics in an SMSA have a negative effect on the labor force participation rate of wives in the area; this negative effect is larger than the positive effect of median wages of employed females. Bowen and Finegan reason that wages of domestics act as an indicator of the relative price of a variety of goods and services used by working wives to substitute for their own time in the home. One of these services is child care. Similarly, Heckman (1974) argues that price of child care is a determinant of labor force participation and of hours worked by employed women. Thus, consistent with past empirical findings concerning the price of child care, our second hypothesis is <u>the greater the cost of child care in a local area, the larger the constraint of children on the market activity of wives who live there</u>. Because the need for child care declines as children grow older, we hypothesize that the impact of child care costs on the constraint of

children is smaller for older children than for younger ones, similar to our reasoning about the relationship between children's ages and the availability of childcare.

Our third hypothesis concerns labor market differences in the availability of jobs which are compatible with the demands of childrearing. Darian (1975; 1976) recently developed the idea that labor force participation of mothers of young children depends on the availability of jobs which combine easily with childrearing responsibilities. Darian classified occupations in which large proportions of incumbents (1) walked to work, (2) worked at home, or (3) worked part time, as "convenient" for mothers of young children. She found that the inhibiting effect of preschool-aged children on their mother's labor market activity dependent somewhat upon the "convenience" of the mother's current or last occupation. In 1960, a higher proportion of white married mothers with children under six were employed if their current or last occupation was "convenient" than if it was not. But the availability of "convenient" jobs is a property of local areas and there is ample reason to suspect that areas differ widely in the "convenience" of the jobs they provide for their residents who must also attend to childrearing. For example, SMSAs differ widely in their occupational composition (see, e.g., Stolzenberg and D'Amico, 1977), and in patterns of journey to work (see U.S. Bureau of the Census, 1972); all three of these factors, as well as others, would seem to have direct impact on the notion of job "convenience" as developed by Darian. So we hypothesize that <u>the constraint of children on their mother's labor force participation is stronger in areas where factors tending to</u>

enhance job "convenience" are present, and that the constraint of children on their mother's labor force participation is weaker in areas where factors tending to enhance job "convenience" are absent. As with our first two hypotheses, we expect that the effect of job "convenience" on the constraint of children will be stronger for younger children than for older children.

We now turn to some methodological issues involved in testing the three hypotheses we have just presented.

Methodological considerations

The hypotheses just presented assert that the relationship between women's family composition and labor force participation probability is partially determined by characteristics of the labor market in which women reside. In more statistical language, these hypotheses assert that if one were to measure the relationship between women's family composition and the probability of labor force participation in each of several different labor markets, the parameters of this relationship would vary from one labor market to another. Further, this variation would be systematically related to other characteristics of these labor markets, such as the supply and price of child care labor. Accordingly, a proper method for testing these hypotheses is to perform separate analyses of women's labor force participation in each of several different labor markets, and then to relate the outcomes of these analyses to the labor market characteristics named in our hypotheses. This is, of course, the same type of research design that underlies the analysis of covariance (see, e.g., Cochran's 1957 classic statement of

this point). The analysis of covariance would be the proper method for testing our hypotheses, were it not for the fact that labor force participation is represented by a dummy variable, and covariance analysis presumes the use of a continuous dependent variable (more on this below). Accordingly, we adopt the general strategy of covariance analysis, but we use methods which are better suited to the dichotomous nature of labor force participation.

As in covariance analysis, our statistical analysis takes place in two steps. The first step is based on individual-level data, while the second step is based on comparisons among individual-level results in different labor markets.

Step 1: Individual-level analyses in each labor market. The problems of using least squares methods such as regression or covariance analysis to analyze the determinants of a dummy dependent variable are well-known and range from frequent uninterpretability of results to bias in statistical tests (see, e.g., Hanushek and Jackson, 1977:179-182; Goldberger, 1964:248-252; Theil, 1971:628-629). Therefore, we use probit analysis, a technique which avoids these problems to estimate an individual-level model of female labor force participation[4] in each of several geographic areas.

[4]Maximum-likelihood logit analysis is another technique which would be suitable for the first-stage and individual-level analyses. Where the effect of a predictor variable is measured by the partial derivative of the probability that a woman is in the labor force with respect to the predictor variable (as is the case in our study), it can be shown that logit and probit analysis yield almost identical results near the means of relevant variables (Hanushek and Jackson, 1977: 188-189). We choose probit analysis over logit analysis because our computer program for probit analysis is more efficient than our program for logit analysis.

The probit model of the first stage analysis takes the form

$$(1) \qquad P = F\left(a + \sum_{i=1}^{k} b_i X_i\right),$$

where:

P is the probability that a given woman participates in the labor force,

F is the cumulative distribution function of the normal distribution,

a is a constant,

$b_1, b_2, \ldots, b_k$ are coefficients, and

$X_1, X_2, \ldots, X_k$ are variables representing characteristics of the woman and her economic, demographic, and social circumstances which affect her probability of labor force participation. (These variables and the reasons for their inclusion are described below.)

The effect of a given causal variable, X_i, on the probability of labor force participation is usually defined as the partial derivative of P with respect to X_i, $\partial P/\partial X_i$, and is a rate of change, expressed as changes in probability per unit change in X_i (see Stolzenberg, 1979). It can be shown that

$$(2) \qquad \partial P/\partial X_i = b_i \, f\left(a + \sum_{i=1}^{k} b_i X_i\right),$$

where f is the normal density function, and all other symbols are as defined above. Since the effect of X_i on the probability of labor force participation is a function of the coefficient for X_i and of the other variables in the model, the effect of X_i is sensitive to the values at which the partial derivative in equation (2) is evaluated. By evaluating the partial derivatives at the same value of P in all geographic areas we avoid confounding areal differences in the determinants of labor force participation with areal differences in the point at which the effect measures are calculated. For a general and lucid treatment of these and related points, see Hanushek and Jackson (1977:Ch. 7).

Step 2: Analysis of labor market impact on individual-level effects. Based on past research and on our own thinking, we have hypothesized that various labor market characteristics intensify or weaken the impact of certain individual-level variables on the probability that women participate in the labor force. We test these hypotheses by relating the effect of an individual-level variable on labor force participation in a given labor market (calculated from the probit analysis described above) to other characteristics of that market. That is, we take the labor market as the unit of analysis in the second-stage analysis, we treat the effect measures from the first-stage analysis as properties of areas, and we regress these effect measures on the labor market characteristics which we hypothesize to intensify or weaken these effects. This second-stage regression analysis will indicate the presence and strength of the hypothesized effects.

Actually, the second-stage regressions are weighted least-squares analyses, to accommodate heteroskedasticity in the effect measures. Transformations of estimated standard errors of the probit coefficients calculated in the first-stage analysis provide necessary indicators of the error variance of the effect measure in each labor market.

We now describe data and variables, and then turn to results.

Data and Results for Step 1

Necessary individual-level data are available from the U.S. Census Bureau's 1970 County Group One-In-a-Hundred Public Use Sample Datafiles. These data contain information from basic census forms, and from the five percent and fifteen percent sample schedules. These data also identify the specific county group (see definition below), in which each respondent resides, thereby permitting separate, individual-level analyses of the female labor force participation in each of these identifiable areas. For present purposes, county groups are ideal geographic areas, since their boundaries follow the geographic outlines of local labor markets. More specifically, the Census Bureau (1972:4-5) describes county groups as follows:

> The 'areas' delineated [by county groups] correspond to economic areas designated by the Bureau of Economic Analysis, Regional Economics Division (or occasionally combinations of related economic areas where necessary to meet population criteria). The areas are based on a nodal-functional area concept. That is, to each urban center are attached those surrounding county units where economic activity is focused directly or indirectly on that center. These areas represent an extension of the SMSA concept--<u>a primary area of economic activity and commuting</u>--<u>to include the rest of the area for which its central city is the trade and labor market center</u>. Thus SMSA's are considered to be integral parts of the larger metropolitan complexes. In rural parts of the country, with

> no SMSA's the comparable economic centers are cities of 20,000 to 50,000 population. [Italics added.]

Several additional features of county groups are important for present purposes: First, there are enough of them (n = 409) to permit a detailed, reliable second-stage statistical analysis. Second, they cover the full range of labor markets in the U.S. Previous studies of labor market effects on female labor force participation (e.g., Bowen and Finegan, 1969; Mincer, 1962, Cain, 1966) have been limited to the largest Standard Metropolitan Statistical Areas (SMSAs) in the country. The largest SMSAs are disproportionately located in the Northeast, and there is little reason to believe that they are representative of labor markets in the nation as a whole. By using county group data, rather than the usual SMSA data, we avoid these problems of nonrepresentative samples.

In the probit analysis of each county group, the dependent variable is a dummy variable for labor force participation. Independent variables include measures of the number and ages of women's children and indicators of other factors which are known to affect married women's labor force participation. Inclusion of these other factors in the individual-level model reduces the possibility that we confound the effects of children with the effects of the correlates of children, and assures that the effects of children are measured net of the effects of other individual and family characteristics which affect the probability that married women work for pay. Control variables include indicators of physical disability which would interfere with employment, years of schooling completed, husband's income, the number of adults in the

household, the woman's age, the age at which she married, the number of times she has married, whether she is nonwhite, whether she is Hispanic, whether the head of her household is retired, and a multiplicative interaction between the woman's age and her husband's income. Schooling is measured with a series of piece-wise continuous variables which allow for nonlinearities in the effect of educational attainment on the probability of labor force participation. The use of piecewise continuous variables allows for more variation in the effect of schooling than does the use of years of schooling and years of schooling squared, especially at pivot points in the educational career, such as 12 years. The piece-wise continuous specification also avoids rigidly parameterizing the effect of education and consumes fewer degrees of freedom than using dummy variables for each year of school. A squared term for years of age allows for nonlinearities in the effect of age on labor force participation.[5]

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[5]Another variable which might be included as a control variable is some measure of the wife's history of prior labor force participation. On the one hand, prior labor force participation is a powerful predictor of current labor force activity (Waite, 1976; Heckman and Willis, 1977), and one might wish to include it in the model to hold constant the wife's long-term propensity toward paid employment. However, the only measure of past labor force activity available in the Census Public Use data is an indicator for the wife's labor force participation in 1965. For some women, this variable would be an appropriate, though crude, measure. But for women who were school-enrolled, or who were simply too young to have been employed five years ago, not having participated in the labor force would not indicate anything about long-term propensities toward paid employment. Convinced that these biases would be significant but not being able to quantify their impact on the use of "activity five years ago" as an indicator of taste for labor force activity, we estimated all of the individual-level models for each county group twice, once including, and once excluding a dummy variable for labor force participation five years ago. We also performed all second stage analyses twice, once with first stage results including the experience variable, and once with first stage results excluding it. Results and conclusions were virtually identical in all cases. For parsimony, we report only the results obtained without the activity five years ago variable.

The constraint of children on labor force participation is measured with six variables: TOTCH indicates the total number of children that the mother has living at home with her. CHO-2 indicates the number of children aged two or less living at home. CH3-4 is the number of children aged three to four. CH5-6 is the number of children aged five to six. CH7-12 is the number of children aged seven to 12 and CH13-18 is the number of children aged 13 to 18. To avoid multicollinearity, there is no indicator for the number of children over 18 living at home. The coefficient for TOTCH indicates the effect of the number of children at home on the probability of labor force participation, but it does not allow this effect to vary with the ages of the children. Coefficients for remaining number-of-children indicators adjust this effect for the differential impacts of children of different ages. Thus, the model allows, but does not require, the constraint of young children to be stronger than the constraint of older children, consistent with our expectations and past research findings discussed above. Table 1 summarizes the names and abbreviations of variables used in the individual-level analyses.

Table 2 summarizes some key findings from the individual-level analyses. That table reports the means and standard deviations of measures of the effect of children on labor force participation. Bear in mind that the the statistics presented in Table 2 are not results of a single probit analysis, but are the means of the coefficients and effect measures obtained in separate analyses performed in each of 409 different geographic areas. The left half of Table 2 presents the means

and standard deviations of the "raw" probit coefficients for the children-at-home variables in the individual-level analyses. The right half of the table presents the means and standard deviations of effect measures computed from the probit coefficients. As discussed earlier, these effect measures are the partial derivatives of the probability of labor force participation with respect to the various children-at-home variables. Each row of the table presents partial derivatives with respect to a different children-at-home measure. To avoid confounding geographic differences in the means of variables with geographic differences in their effects on labor force participaion probability, the partial derivatives are evaluated at the same point in all county groups. Since about 30 percent of the married women with children under six years of age participated in the labor force in 1970, the year to which our data refer, we evaluate the partials at this point. These partials are rates of change in the probability of labor force participation per additional child, measured at a single point on the S-shaped curve relating independent variables in our probit model to the probability of labor force particpation. According to the probit model, number of children is measured as a continuous variable. But children are aquired only in integer increments, and so the effect of an additional child would be obtained by integrating the partial over the domain from, say, zero children to one child, at given values of other variables in the probit model. The probit model assures that this integration does not push probabilities above unity or below zero. As Stolzenberg (1979) points out in a general treatment of effects in nonlinear models, the integral is cumbersome to work with, and the

partial derivative provides a useful, mathematically meaningful index of effect, even though it is expressed as a rate of change in probabilities rather than actual changes in probabilities.

Looking at the means of the partials (column "C"), it is clear that the mean of effects found in our analysis are entirely consistent with past findings based on data which aggregate all areas of the U.S. together: The presence of children under two years old exerts a powerful inhibiting effect on participation. This effect decreases as the ages of children increases, until children over 13 years of age show a small positive effect on labor force participation.

However, our main concern here is with geographic variation in the effect of children on labor force participaion. Table 2 shows that variation to be considerable. For example, looking at columns "C" and "D" in the row for CHO-2, notice that the effect of each child aged two or less is, on the average, -.7786. That is, the average effect of an additional child is to reduce the probability of labor force participation at a rate of 77.86 percentage points per additional child. However, notice that the standard deviation of the effect of CHO-2 is .2250. Thus, at one standard deviation above the mean the effect of an additional child on labor force participation probability is -.5536, while the effect at one standard deviation below the mean is -1.0036 (-.7786 -.2250 = -1.0036; -.7786 +.2250 = -.5536). Thus, at one standard deviation below the mean, the effect of CHO-2 on probability of participation is about 81 per cent larger than the effect at one standard deviation above the mean (-1.0036 / -.5536 = 1.81). This variation in the effect of children on labor force participation seems

large indeed, and Table 2 makes clear that the impact of children in other age brackets on their mothers' labor force participation probability also varies considerably across county groups. This finding is consistent with the reasoning that motivated us to undertake this research. We now turn to data for testing our ideas about the causes of this variation.

Data and Results for Step 2: Labor Market Impact on Individual-Level Effects

To test our hypotheses about the effects of childcare availability and cost, we require measures of the cost and availability of childcare in county groups. Ideally, one would measure the availability and cost of childcare with direct measures of the cost of childcare services and the difficulty which parents of small children encounter in finding suitable care for their offspring. However, such data do not exist, except for a few areas or at the national level (see Hofferth, 1979), and so we resort to indicators which are based on Census Public Use data. Our measures of childcare cost and availability are based upon the labor intensive nature of this activity. Although there is some variation in the number of children that an adult can supervise, the realities of life with children place a fairly low upper bound on the ratio of children to adults. (The upper bound on the child-adult ratio varies with the age of children, but our analyses allow the effects of children to vary with their ages.) Further, laws in many states limit the ratio of children to child care workers in care facilities. Accordingly, we use the number of persons in an area who report paid employment as child care workers as a measure of the capacity of child

care facilities in that area.[6]

By "availability of childcare" we mean the extent to which child care capacity corresponds to the number of women who might require childcare to permit them to work for pay. We measure this correspondence with the ratio of childcare workers per female labor force participant in the area. But because childcare services are used both by women who work and by women who do not work, and because some women with children might be out of the labor force because of difficulty in locating childcare, we also measure childcare availability with the ratio of the number of childcare workers to the number of women who are "eligible" for labor force participation. By "eligible," we mean women who are between the ages of 16 and 65, who are not enrolled in school, and who are not members of the armed forces. By "child care workers" we mean persons whose occupations are classified by the Census as "child care workers outside private households" or "child care workers inside private households." To reduce skewness in their distributions, we take natural logarithms of both measures of availability.

To measure the cost of childcare in a county group, we use the mean natural log implied hourly earnings of childcare workers who are employed outside of households in that area. We reason that the labor intensive nature of childcare makes the hourly earnings of childcare workers a serviceable index of the cost of childcare services. We exclude private household workers from these calculations because of

[6]As used here, "facilities" means childcare which takes places in or out of the child's own home, so long as the adult providing the services is doing so for pay.

difficulties in rendering private household workers' pay comparable to the pay of other workers.[7]

To measure the "convenience" of jobs for working wives in county groups, we use the 1972 City and County Data Book (Bureau of the Census, 1972) to construct two variables. Following Darian's original conceptualization of "convenience," both measures focus on the journey from home to job. Our first measure is a crude index of the average distance from home to work. Hereafter called WORKOUTC, it is the percent of employed persons in the county group whose jobs are located outside the county in which they reside. Crude though it may be, this measure is based on the fact that travelling longer distances tends to involve more crossing of county lines than travelling shorter distances, other things being equal. The variation in county size adds error to this measure, but information on journey to work is exceedingly limited and we think it better to utilize crude measures than to use no data at all on this aspect of employment. Following our hypotheses, we expect that the constraint of children is larger in areas with large values of WORKOUTC than in areas where WORKOUTC is smaller.

The second "convenience" measure used here focuses on the mode of transportation rather than the distance travelled. While private

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[7]Following the standard procedure for estimating hourly wages from Census public use data, we estimate mean log implied hourly earnings as the mean of the natural logarithms of implied hourly earnings, where implied hourly earnings is earnings in 1969 divided by the product of weeks worked in 1969 and hours worked in the Census week. The strengths and weaknesses of this procedure are well known and are reviewed in Stolzenberg (1975). Among the difficulties in comparing wages of private household workers to other workers are problems in evaluating payments in kind, such as meals and lodging, and problems in comparing reported income, on which employment and income taxes are paid, and unreported income, on which no taxes are paid.

transportation adjusts to the needs of its riders, departing at times they select and travelling by routes of their choice without delays at intermediate stops, public transportation normally does not do so. Thus, public transportation is less convenient than private transportation, both in the sense of Darian, and in the more general meaning of the word. Accordingly, we use the percent of employed persons in a county group who travel to work via public transportation as a measure of the convenience of jobs in the county group. We reason that the smaller the proportion using public transportation to get to work, the more "convenient" the jobs, on the average. We label the percent of employed persons who travel to work via public transportation PUBTRANS. In passing, we mention that the crudeness of our "convenience" measures is apparent to us, and that our second stage empirical analyses are performed both with and without PUBTRANS and WORKOUTC, to allay any fears that errors in these variables contaminate findings regarding the relationship between other county group characteristics and the constraint of children on their mothers' labor force participation. We now briefly describe the distributions of the county group variables, and then move on to examine their effects on the constraint of children measures obtained from the first stage analyses.[8]

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[8]There is considerable awareness in sociology of problems of interpretation caused by the use of regressors which are ratios or differences of other variables (see, e.g., Fuguitt and Lieberson, 1974; Schuessler, 1973). Our measures of childcare availability are logarithms of such ratios, and so are differences between logarithms of their numerators and denominators. To investigate the possibility that we were confounding the effect of the denominator with the effect of the ratio, we calculated all of the second step regressions reported in this paper with the addition of the log of the denominator of the childcare availability ratio as an independent variable. These regression analyses did not lead to different substantive conclusions from the regressions reported in Tables 4 and 5, nor did they differ substantially in

There are two important observations to make about the county group characteristics described above. The first is that, at the empirical level, it is not necessary to distinguish between the number of childcare workers per female labor force participant and the number of childcare workers per woman who is "eligible" for labor force participation: the logarithms of these two variables correlate +0.98 with each other, making them near perfect substitutes for each other. We report results only for the former measure. The second point is that all five of the county group characteristics vary substantially. Table 3 presents the means and standard deviations of these variables. Because the childcare cost and availability measures are measured on logarithmic scales, their means and standard deviations must be exponentiated for intuitively meaningful interpretation. The exponentiated mean of the logarithm of child care workers per female labor force participant is 0.0240, or about two and one half child care workers per hundred working women. At one standard deviation above the mean, the exponentiated value is 4.28 child care workers per hundred women labor force participants; at one standard deviation below the mean, the exponentiated value is 1.35 per hundred. The value at one standard deviation above the mean is 3.18 times as large as the value at one standard deviation below the mean. Clearly, there is substantial variation across county groups in this measure of child care availability. Similar results obtain for the number of childcare

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any other way. But, since we have no substantive reason for including the logs of the denominators of these variables in our models, we do not present the regressions which included the logs of the denominators of the availability measures.

workers per woman who is "eligible" to participate in the labor force. Similarly, there is considerable variation in the cost of child care. Exponentiating the mean of child care cost (0.2393) yields a value of only $1.27 per hour. But the standard deviation of this variable is 0.6582, which indicates that at one standard deviation above the mean, the cost of childcare is almost twice the value at the mean; the value at one standard deviation above the mean is about three and three quarters as large as it is at one standard deviation below the mean.[9] Notice that the standard deviation of PUBTRANS is larger than its mean, and that the standard deviation of WORKOUTC (12.6 percent) is quite substantial, particularly when compared with the mean of 17.9 per cent. So there is substantial variation across county groups in the variables which we have argued affect the constraint which a woman's children put on her labor force participation.

To test our hypotheses that the constraint of children depends on local area characteristics, we regress the probit coefficients for the number and ages of children at home on measures of the cost and availability of childcare.[10] Table 4 gives results when the constraints of children are regressed only on cost and availability of childcare measures; Table 5 gives findings obtained when probit

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[9]Exp(0.6582) = 1.93, which is almost two. Exp (2 x 0.6582) = 3.73, which is about three and three quarters.

[10]As noted earlier, these regressions are weighted least squares computations, in which case weights are proportional to the inverses of the squared standard errors of the probit coefficients. Standard errors are obtained from the probit analyses which generated the coefficients. Our use of probit coefficients rather than effect measures evaluated at some fixed point is of little concern here, since the effect measure evaluated at, say, the 0.3 probability level is a constant multiple of the probit coefficient.

coefficients are regressed on childcare measures and job convenience indicators. Note that each of these tables reports the results of six different regression analyses. Each line of these tables reports a regression in which the dependent variable is the probit coefficient for a different number-and-age-of-children variable. Thus, the first line of Table 4 reports an analysis in which the probit coefficient for CH0-2 (number of children at home 0 to 2 years old) is the dependent variable. The second line reports an analysis in which the dependent variable is the probit coefficient for CH3-4, the number of children at home aged three to four. The unit of analysis in these regressions is the county group.

Recall that we expected the constraint of children on women's labor force participation in an area to vary directly with the cost, and inversely with the availability, of childcare there. Since the constraint of children is a negative effect of children on the probability of labor force participation, our hypotheses suggest positive effects of the availability of childcare, and negative effects of the cost of childcare, on the probit coefficients for the number-and-age-of-children variables. However, we also expected that these relationships would be stronger for younger children and weaker for older offspring. Looking at column "A" of Table 4, notice that findings for the availability of childcare correspond precisely to expectations: In the row for the analysis of the probit coefficient for CH0-2, the large positive and statistically significant effect of child care availability indicates that the constraint of children younger than two years is strongly affected by the availability of childcare. The

standardized coefficient of 0.37 is substantial by normal standards, and the metric coefficient indicates substantial sensitivity of the probit coefficient for CHO-2 to the availability of childcare.

Looking at the next row of Table 4, notice that the effect of childcare availability on the probit coefficient for children aged three to four is also large, positive and significant, but that it is smaller than the effect of childcare availability on the constraint of children aged two or less. Similarly, the effect of childcare availability on the probit coefficients for number and age of children declines monotonically as the age of the children increases. Notice that the effects of childcare availability on TOTCH and CH13-18 are trivial. These findings are entirely consistent with the hypotheses presented earlier in this paper.

Now look at columns "C" and "D," to see the effect of child care costs in a county group on the relationship between children and their mother's probability of labor force participation in that area. As in the case of childcare availability, the coefficients for child care cost has the expected sign, and the effect is statistically significant for the probit coefficient for CHO-2. But child care costs has no statistically or substantively significant effect on the constraint of children older than two years. This finding is consistent with our hypothesis, but it suggests that availability, rather than cost, drives the relationship between childcare and the constraint of children in an area, and that the cost of childcare raises the constraint of very young children, but does little to affect the impact of older children on their mothers' labor force participation.

In Table 5, we add measures of job "convenience" in county groups to the area characteristics considered in Table 4. Before looking at the effects of PUBTRANS and WORKOUTC, note that adding convenience measures has done nothing to change the conclusions we drew from Table 4. Comparing the two tables, notice virtually the same pattern of effects for childcare cost and availability measures, except that effects in Table 5 are slightly smaller than in Table 4. This decrease is to be expected when additional control variables are added to a multiple regression equation. Turning now to the effects of job "convenience" on the constraint of children, recall our hypothesis that the constraint of children increases with both the proportion of workers who are employed outside of their county of residence, and the proportion of workers who travel to their jobs by public transportation. Since the constraint of children is a negative effect of children on the probability of labor force participation, our hypotheses suggest negative effects of PUBTRANS and WORKOUTC in the regression equations presented in Table 5. Looking at columns "G" and "H" in Table 5, notice that the effect of PUBTRANS on the constraint of children is statistically significant for all age groups (t statistics larger than 2.0 in the first five rows of Table 5). The coefficients for PUBTRANS all have the expected sign too. And while the unstandardized coefficient for PUBTRANS is the same in the regressions pertaining to CH0-2, CH3-4 and CH5-6, the coefficient is smaller for CH7-12 than for younger ages, and smaller still for CH13-18. These findings are consistent with our hypotheses, though the impact of PUBTRANS on the constraint of children is modest, even at its strongest.

Looking at columns "E" and "F" notice that the effect of commuting distances, as measured by the proportion of employed persons who work outside their county of residence, is approximately what we hypothesized: In the analysis of the probit coefficient for CH0-2, the effect of WORKOUTC is negative and statistically significant (t = -2.841). Looking at the next line in Table 5, notice that the effect of WORKOUTC on the probit coefficient for CH3-4 is also statistically significant (t = -4.046) and negative, though it is slightly larger than in the analysis of the probit coefficient for CH0-2. WORKOUTC has no statisically significant effects in other analyses reported in Table 5, though the signs of these insignificant coefficients are consistent with our hypothesis that job "convenience" in an area affects the impact of children on the labor force participation of mothers who work there. However, the results for WORKOUTC lack the neat relationship with age of children that we observed with the childcare cost and availability measures, and it is clear that WORKOUTC has only modest impact on the constraint of children on their mother's labor force participation.

Discussion and Conclusions

At the start of this paper we hypothesized that the well-known inhibiting effect of children on their mother's probability of labor force participation is shaped by characteristics of the place in which the woman lives. Our hypotheses were based on the observation that incompatibility between children and their mother's employment is determined by the characteristics of the jobs available to the mother, by the availability and cost of services which working mothers utilize

to deal with these incompatibilities, and by the age-varying demands which children place on those who must care for them. To test these hypotheses we drew upon vast amounts of census data and employed research methods which overcame most of the problems which have plagued analyses of dichotomous dependent variables until recently. These data and methods have enabled us to examine the relationship between individual and aggregate determinants of labor force participation in unusually fine detail, free from some of the methodological problems which have at times cast doubt on previous research findings on women's work activity. In general our findings are very consistent with our hypotheses.

Several additional points about our findings bear emphasis. First, we note that the ages of a woman's children, and not just their numbers, are critical to understanding the constraint they place on their mother's labor force activity. Not only did we find it essential to allow for children's age differentials in constraints in the individual-level probit equations (which we expected from the findings of prior research, e.g. Sweet, 1973), but we also found that the relationship between area characteristics and the constraints of children was stronger for the constraints of younger children than for the constraints of older offspring. In short, our analyses had much to say about the relationship between area characteristics and the constraints of young children, but they were less informative about the effect of area characteristics on the constraint of older children.[11]

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[11]To quantify the explanatory power of our analyses of the constraint of young children, we square the multiple correlation coefficient for the second-step equations explaining variation in the probit coefficient for CHO-2. In the analysis reported in Table 4, the R-

Further thought and research is needed to understand the constraint of older children.

Another point which bears mentioning is the absence of attitudinal factors in our second step models. These factors are excluded because of data limitations, not because of our theoretical perspective. We suspect that prevailing attitudes in an area do indeed shape the way that mothers and fathers respond to the demands of childrearing, even if prevailing attitudes are different from their own opinions. However, investigation of the effect of community attitudes on the constraint of children must await development of adequate data.

We also think it important to raise the usual cautionary concerns about imputation of causal ordering from cross-sectional data. We have argued that the cost and availability of childcare, and the "convenience" of jobs for mothers of young children affects the constraint of children on their mother's labor force participation. Our findings are consistent with our argument. But one might argue that causality proceeds in the opposite direction, and that jobs are made convenient, and that childcare is made cheap and plentiful, by low constraints of children. Though we find these arguments implausible, the data we have examined would be consistent with them too. Without longitudinal data, we have no empirical basis for rejecting different imputations of causal direction than have been made in this paper, however implausible they may seem.

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squared statistic is 0.1501, indicating that the cost and availability of childcare variables explain 15 percent of the variance in the coefficient for CHO-2. In the corresponding analysis from Table 5, 22 percent of the variance in the same coefficient is explained. However, we hesitate to make much of variance explained measures (see Duncan, 1970; Stolzenberg and Land, 1981), and so we raise this point only in passing.

The caveats and limitations of our analyses notwithstanding, we think that the results described here are remarkably consistent with the general perspective outlined at the start of this paper. That perspective, which we believe is quintessentially sociological, stresses the role of aggregate-level phenomena in shaping the behavior of individuals. Our analyses show that this shaping involves substantively and statistically significant effects of local areas on the impact of those most powerful constraints on a mother's employment, her children. We are hopeful that our findings will stimulate further study of the relationship between aggregate and individual level determinants of women's economic activity. Such studies appear to hold considerable promise.

Table 1--VARIABLES FOR INDIVIDUAL-LEVEL MODEL FROM PUS 1/100 (5% SAMPLE)

Variable Name	Description
TOT CH	Number total children
CH 0-2	Number of children 0 to 2 years old.
CH 3-4	Number of children 3 to 4 years old.
CH 5-6	Number of children 5 to 6 years old.
CH 7-12	Number of children 7 to 12 years old.
CH 13-18	Number of children 13 to 18 years old.
CH 19+	Number of children 19 or older
DLTY	Wife disability--Does the wife have a disability which limits the amount of kind of work or prevents any work; 1 = yes, 0 = no.
DBDW	Length of wife disability in years
ED08	Years of schooling if years of schooling = 0-8 years; 0 else
ED911	Years of schooling if years of schooling = 9-11 years; 0 else
ED12	Years of schooling if years of schooling = 12 years; 0 else
ED13	Years of schooling if years of schooling = 13-15 years; 0 else
ED16	Years of schooling if years of schooling = 16 years; 0 else
ED17	Years of schooling if years of schooling $\geq$ 17 years; 0 else
HUBI	Family income less wife earnings
ADUL	Number adults in household
AGE	Age of wife in years
AGEH	Age of household head in years
AGES	Age of household head, squared
AGEM	Wife age at first marriage, in years
TSMA	Wife married more than once; 1 = yes, 0 = no
NONW	Nonwhite; 1 = yes, 0 = no
SPAN	Spanish surname; 1 = yes, 0 = no
HDRT	Husband disability, husband retired or disabled; 1 = yes, 0 = no
AGHI	Age * FILOW interaction

Table 2

MEANS AND STANDARD DEVIATIONS OF PROBIT COEFFICIENTS FOR NUMBER AND AGES OF CHILDREN

	Raw Coefficients		Effects Evaluated at P=30.0%	
	Mean	Std. Dev.	Mean	Std. Dev.
Variable	"A"	"B"	"C"	"D"
TOTCH	.0175	.0489	.0061	.0170
CH0-2	-2.2389	.6472	-.7786	.2250
CH3-4	-1.3188	.6345	-.4586	.2207
CH5-6	-.9238	.5274	-.3213	.1834
CH7-12	-.6415	.4134	-.2231	.1438
CH13-18	.0154	.4049	.0054	.1408

Table 3

MEAN AND STANDARD DEVIATION OF COUNTY GROUP CHARACTERISTICS USED IN SECOND STEP ANALYSIS

Variable	Mean	Standard Deviation
1. Child Care Availability		
a. Ln (Child care workers per female labor force participant)	-3.7299	.5787
b. Ln (Child care workers per "eligible" women)	-4.3430	.5658
2. Child Care Cost (Ln (dollars/hour))	.2393	.6582
3. PUBTRANS (%)	4.7094	7.9033
4. WORKOUTC (%)	17.9268	12.6122

Notes: See text for variable descriptions.

Table 4

WEIGHTED LEAST SQUARES REGRESSIONS OF CONSTRAINT OF CHILDREN ON COST AND AVAILABILITY OF CHILDCARE (CHILDCARE WORKERS PER FEMALE LABOR FORCE PARTICIPANT)

Dependent Variable and Probit Model from which Estimated:	Ln (Child-Care Workers Per Female Labor Force Participant)		Mean Ln (Implied Hourly Earnings of Child-Care Workers)		
	b/t	B	b/t	B	R
CH0-2	.381 (8.128)	.37	-.130 (-2.914)	-.13	.4154
CH3-4	.265 (5.497)	.27	-.033 (.738)	-.04	.2734
CH5-6	.213 (5.455)	.26	-.046 (1.235)	-.06	.2788
CH7-12	.156 (4.965)	.24	-.055 (1.800)	-.09	.2712
CH13-18	.041 (1.406)	.07	-.031 (-1.085)	-.05	.0967
TOTCH	-.010 (-2.741)	-.14	.001 (.404)	.02	.1408

Notes: N = 408 county groups.
See text for full definitions of variables.

Table 5

WEIGHTED LEAST SQUARES REGRESSION OF CONSTRAINT OF CHILDREN ON COST AND AVAILABILITY OF CHILDCARE AND MEASURES OF JOB CONVENIENCE

Dependent Variable and Probit Model from which Estimated:	Ln (Child-Care Workers Per Female Labor Force Participant)		Mean Ln (Implied Hourly Earnings of Child-Care Workers)		WORKOUTC		PUBTRANS		
	b/t	B	b/t	B	b/t	B	b/t	B	R
	"A"	"B"	"C"	"D"	"E"	"F"	"G"	"H"	"I"
CH0-2	.257 (4.777)	.250	-.102 (-2.287)	-.104	-.005 (-2.841)	-.129	-.010 (-3.532)	-.189	.4687
CH3-4	.127 (2.324)	.128	-.002 (-.053)	-.003	-.008 (-4.046)	-.194	-.010 (-3.299)	-.185	.3724
CH5-6	.111 (2.426)	.137	-.018 (-.475)	-.023	-.001 (-.722)	-.035	-.010 (-3.990)	-.229	.3403
CH7-12	.064 (1.762)	.100	-.025 (-.814)	-.040	-.002 (-1.478)	-.073	-.007 (-3.956)	-.231	.3396
CH13-18	.011 (.301)	.018	-.023 (-.763)	-.039	.002 (1.231)	.064	-.004 (-2.069)	-.128	.1483
TOTCH	-.006 (-1.402)	-.084	.000 (.096)	-.005	.000 (-.221)	-.011	.000 (1.679)	.102	.1631

Note: t-statistics given below coefficients.

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